Collective Agreement Status and Survivability: Change and Persistence in the German Model*

John T. Addison
University of South Carolina, GEMF, and IZA
Paulino Teixeira
University of Coimbra, GEMF, and IZA
Alex Bryson
National Institute of Economic and Social Research and CEP
André Pahnke
Institut für Mittelstandsforschung Bonn

Abstract

This paper assesses the decline in collective bargaining coverage in Germany. Using repeat cross-section and longitudinal data from the IAB Establishment Panel, it indicates the overwhelming importance of behavioral as opposed to compositional change in this process. Further, in the first use of survival analysis for the purpose, it also charts workplace transitions into and out of collective bargaining. In addition to providing new estimates of the median duration of coverage, the paper reports on the factors generating entry into and exit from collective bargaining. These influences are found to be distinct but symmetric.

JEL Classification: J52, J53

Keywords: collective agreement coverage, bargaining transitions, survivability, Germany

* The authors wish to thank the Editor of this journal and two anonymous reviewers for their constructive suggestions. Addison gratefully acknowledges financial support from the *Riegel and Emory Foundation* of the University of South Carolina.

I. Introduction

Use of the characteristics of collective bargaining to help motivate analysis of wage and employment outcomes occupies an important position in contemporary treatments of the covariation of institutions and macroeconomic outcomes. Thus, notions of the importance of the centralization of collective bargaining (or its absence) to wage and unemployment development figured heavily in policy discussions in the 1980s (Calmfors and Driffill, 1993). More recently, the importance of centralization has been supplemented if not supplanted by notions of coordination (e.g. OECD, 2004, Chapter 3). *Vulgo*: more centralized bargaining regimes – latterly, more coordinated ones – have been held out as offering scope for improved economic performance.

One important issue that has arisen is the stability of the bargaining structure (and of that of labour institutions more generally) through time. At best, bargaining structures – centralized, decentralized, and/or coordinated – have been observed at a few points in time, and insufficient attention has been paid to within-country changes. This seems singularly inappropriate when industrial relations systems are evidently under sustained stress. At another level, most studies investigating time-varying changes in collective bargaining have with a few notable exceptions (e.g. Dustman, Ludsteck, and Schönberg, 2009) focused on union density. This strategy also seems inadequate given that union density and collective bargaining coverage are far from being perfectly correlated.

In the present study we will focus on Germany, the exemplar of moderately centralized but well-coordinated industrial relations in continental Europe, and the interval 2000-2008. Despite its long-standing reputation as a beacon of stability, however, the German case is also of particular interest given the recent signs of atrophy in its industrial relations architecture. Here we refer of course to the erosion of sectoral (i.e. industry-wide) collective bargaining.

The four component parts of this paper should be seen as proceeding in tandem. First, we offer descriptive information on the hemorrhaging of collective bargaining. This material not only plays an important part in scene setting but also hints at the potential importance of certain variables (viz. establishment size and region) while providing some practical justification for our subsequent concatenation of the categories of sectoral and firm-level bargaining into a single "any collective bargaining composite." Second, a decomposition exercise, based on two cross sections of the data not only confirms an earlier result in the literature that the decline in collective bargaining cannot be attributed to changes in the make-up of the workforce or in establishment characteristics (see, in particular, Addison et al. 2011), but also introduces the full set of arguments to be used in the rest of our analysis. Importantly, it also establishes the case for treating establishment heterogeneity more directly. Third, our random effects treatment of collective bargaining propensity is grounded in this recognition, exploiting a longitudinal panel to explain why some establishments may be expected to be covered by collective agreements while others are likely to populate a zone free of collective bargaining. Besides addressing the role of the covariates, it also quantifies the inertia in collective bargaining status. Finally, we turn to the factors that induce failure, examining transitions into or out of a collective agreement. Our survival modeling of collective bargaining takes the analysis of inertia and contract determination one crucial stage further. It is at once the major contribution of the paper and culmination of a lock-step treatment of change and persistence in the German model.

The novelty of our approach stems from the fact that the panel we analyze is long enough to follow some establishments from the outset (i.e. from their birth) to the point of 'failure', here the year in which they switch collective bargaining regime. For these cases, we provide the first estimates of the median duration of coverage, as well as the main determinants of the decision to leave/join a collective agreement. A second innovation involves an analysis of the duration of coverage of panel stayers via the use of a simulated counterfactual. Panel stayers are all those establishments that are left-censored in 2000 (i.e. those with unknown elapsed time of coverage in the year the establishment is first observed).

The use of a nine year observation window in this study is a significant departure from standard cross-section analyses and from panel studies that rely on two distinct moments in time (consecutive or not). Our approach proves fruitful in the sense that it allows us to tackle the determinants of coverage (or absence of coverage) without assuming away unobserved establishment heterogeneity, while at the same time yielding estimates of the duration of different collective bargaining regimes. In the former case, we can ultimately estimate the odds of being covered (uncovered) by collective bargaining at different moments in time. In the latter case, we can model the probability that a given establishment will switch out of (or into) a collective agreement given that it was covered (uncovered) up to certain point in time. Both aspects assist us in interpreting the ongoing decline in collective bargaining coverage in Germany.

II. Preliminaries: The Decline in Collective Bargaining and Worker Representation

Recent studies have documented a shift in the locus of wage bargaining from industry-wide sectoral bargaining to lower levels, both firm-level collective agreements and, much more importantly of late, individual bargains or contracts between the worker and the firm (see Kohaut and Schnabel, 2003a,b; Addison et al., 2010, 2011; Ellguth and Kohaut, 2010, 2011). There has also occurred a decentralization *within* sectoral agreements themselves, which development has stimulated a distinct debate as to whether the process is internally destabilizing or not (see, inter al., Hassel, 1999, 2002; Klikauer, 2002; Frege, 2003; Doellgast and Greer, 2007; Haipeter and Lehndorff, 2009; Bispinnck et al., 2010). In any event, the decentralization of sectoral bargaining – occasioned by opening clauses and pacts for competitiveness (see, respectively, Schnabel and Kohaut, 2007; Seifert and Massa-Wirth, 2005) – has inevitably narrowed the distinction between sectoral and firm-level collective agreements and may also have militated against further growth in the latter. Partly for this reason, once we have charted the course of all types of agreement in the descriptive part of the present exercise, we will subsequently collapse sectoral and firm-level collective bargaining into a single "any collective bargaining" composite for analytical purposes. But in so doing we do not wish to understate the importance of external erosion, namely the sustained increase in the percentage of establishments and employees not covered by any type of collective agreement.

Since the data used in the present exercise differ in a small number of respects from those used in other studies,[1] it is first necessary to provide information on this process of change in collective bargaining. (The Data Appendix succinctly describes our database.) The main trends over the years 2000-2008 are charted in Table A.1 by employment and establishment coverage. The most notable features of the reported values are the decrease in sectoral bargaining and the rise of the collective bargaining free zone. Thus, the share of workers (establishments) covered by sectoral agreements has fallen by 9.5 (11.9) percentage points to 49.2 (36.4) percent, while the share of workers (establishments) without any collective bargaining has correspondingly grown by 8.5 (11.9) percentage points to 42.8 (60.8) percent. The decline in sectoral bargaining has occurred in both eastern and western Germany, even if the rate of decline in bargaining coverage has been more pronounced in the latter region. Interestingly, there has been very little change in firm-level bargaining in terms of establishment coverage, even if employment coverage has risen modestly in both parts of Germany and is higher in eastern Germany.

Coverage rates by establishment size are contained in Table A.2. First, *levels* of sectoral bargaining are considerably higher in larger establishments (defined as those with 250 or more employees), where the collective bargaining free zone is also correspondingly smaller. Larger plants are also more likely to have firm-level collective agreements than their smaller counterparts, although the disparities here are very much smaller. In terms of trends, both size classes register declines in sectoral bargaining (proportionally larger among smaller establishments) and growth in absence of collective bargaining (somewhat greater among larger plants). Modest increases in firm-level bargaining are recorded for larger firms and even for smaller firms based on the employment coverage indicator, but just as in the aggregate case firm-level bargaining is dominated by sectoral bargaining.

A further breakdown of the data by region reveals some differences, while confirming the broad differences between plants of different size and trends in sectoral agreements and no collective agreements. First, absence of collective bargaining is and has remained more important in eastern Germany even if the growth rates are proportionately smaller than in western Germany, especially among smaller establishments. Second, and conversely, sectoral bargaining coverage is lower among smaller and larger establishments in eastern Germany, although the decline has now been proportionately greater among larger east German establishments. Finally, firm-level bargaining is roughly twice as important in eastern Germany for both sizes of establishment than in westerm Germany and any recorded growth is also higher in the east although confined to larger establishments.

Next, if we take any two distinct cross-sections – say 2000 and 2008 – it is analytically straightforward to decompose the decline in collective bargaining coverage into its Oaxaca-Blinder components, namely the so-called 'between' or compositional effect and the 'within' or behavioural effect. The between effect, or the *explained component*, is that part of the observed change that can be attributed to differences in observable characteristics. The within effect, or *unexplained component*, measures the change in coverage arising from differences in propensities (or coefficients). More formally, let $x_{2008}b_{2000}$ be the 2008 (predicted) coverage based on year 2000 coefficients, where x denotes the mean vector of observed (establishment) characteristics and b indicates the vector of estimated coefficients in the corresponding year. The between effect is then given by $(x_{2008} - x_{2000})b_{2000}$ and the within effect by $x_{2008}(b_{2008} - b_{2000})$, where the reference groups are the year 2000 coefficients and the year 2008 characteristics, respectively.[2]

Our selected vector of covariates x includes establishment size, the proportion of skilled and female workers, and dummies for single-establishment status, foreign ownership, establishment age, state of technology, industry, and region (see the Data Appendix for a fuller description as well as sections III and IV below). The broad rationale for their inclusion is to be found in for example Willman, Bryson, and Gomez's (2007) modeling of employer voice-choice decisions. Based on the argument that firms face non-trivial switching costs (i.e. costs connected with uncertainty surrounding the benefits from moving from coverage to non-coverage, and vice-versa), one would expect the returns to being covered by collective agreements to be higher in large establishments and in plants integrated in multi-site establishments. Establishments with a higher proportion of low-skill employees are also likely to rely less on voice mechanisms and therefore expected to be associated with a lower presence of collective agreements. By the same token, older establishments are more likely to be covered given that the incidence of collective bargaining tended to be higher in the past. Alternatively, recently established businesses often require more flexible institutional arrangements at early stages of their existence (see Schnabel et al., 2006).

The results for the entire sample period for collective bargaining of any type are provided in panel (a) of Table A.3. Consistent with prior research, there is clear confirmation that the within effect is overwhelmingly dominant, accounting for more than 90 percent of the observed change in coverage. Interestingly, the dominance of the within effect is no less pronounced in eastern Germany where the decline in overall collective bargaining has been less acute than in western Germany. Panels (b) and (c) of the table repeat the decomposition exercise for the two sub-periods, 2000-2004 and 2005-2008, respectively. (The unemployment peak of 11.7 percent in 2005 provides the basis for the cutoff year.) As it is apparent, the within effect is also clearly dominant in both sub-periods.

It is the precisely this limited role of *observables* found in decompositions based on the 2000 and 2008 cross-sections that sets the stage for our subsequent longitudinal analysis of the propensity of a given establishment to be covered. Essentially, the large within effect we have estimated implies that one has to treat establishment heterogeneity more directly. With these preliminaries behind us, we now turn to a closer examination of this issue.

III. Collective Bargaining Coverage Propensity

Let Y_{it} represent the coverage outcome for the t^{th} observation in the i^{th} establishment. Given the random effect u_i which represents the establishment's persistent *unobserved traits*, namely its unobserved propensity to be covered by a collective agreement, the random-effects probit model can be specified as

$$\Pr(Y_{it} = 1 | u_i, X_{it}) = \phi(X_{it}\beta + u_i\,) \,, \qquad (1)$$

where ϕ is the standard cumulative distribution function and $u_i | X_{it} \sim N(0, \sigma_u^2)$. Model assumptions are therefore that u_i and X_{it} are independent and that u_i has a normal distribution.[3] X includes all observed establishment characteristics that have an impact on the binary response probability, and β denotes the set of parameters to be estimated.[4] As before, the arguments are establishment size, the proportion of skilled and female workers, and dummies for single-establishment status, foreign ownership, establishment age, state of technology, industry, and region. A full set of time (year) dummies is also included in the regression model**.**

Conditional on (u_i, X_{it}), outcomes Y_{i1}, Y_{i2}, …, Y_{iT} are independent, with probabilities depending on u_i and X_{it}. This means that, conditioning only on X_{it}, Y_{i1}, Y_{i2}, …, Y_{iT} will be dependent across t. A useful statistic therefore is the (latent) intra-class (establishment) correlation, given by $\rho = \sigma_u^2/(\sigma_u^2 + 1)$, which indicates the relative importance of the unobserved effect u_i or the correlation between any two observations in the same establishment (see, for example, Rodríguez and Elo, 2003). We will also exploit an additional measure of (manifest) association based on the actual binary outcomes Y_{it}, rather than on the latent variable Y_{it}^*, namely Pearson's r coefficient. The two measures – ρ and Pearson's r – will therefore give the extent to which there is inertia in an establishment's collective bargaining status. Along with these measures, we will also use other statistics (on which more below) evaluated with the linear predictor set at various percentiles, the goal being to have different measures of the persistence of collective bargaining status.

The analysis of collective bargaining coverage propensity by type of collective agreement is presented in Table 1. We retain in the sample all plants surveyed in the 2000-2008 observation window, including those that changed their collective bargaining status more than once. As a practical matter, however, dropping the latter establishments produced virtually no change in the results. From Table 1 it can be seen that establishment size and establishment age are positively and single-establishment firm status negatively associated with coverage. This propensity is also increasing in the skill composition of the workforce. With a few exceptions, the industry and region dummies are also statistically significant. However, other than the lower propensity of establishments located in east German *Länder* to be covered by a collective agreement (not shown in the table), there are no obvious patterns in the data in this regard.

(Table 1 near here)

Running our random-effects probit separately for the two halves of Germany indicates the consistent influences of establishment size and age, proportion of skilled workers, and single-establishment firm status on propensity to be covered. However, the role of foreign ownership in promoting coverage clearly reflects the situation in eastern Germany, while the negative but statistically insignificant influence of workforce proportion female found for Germany as a whole masks the marginally significant effect of that argument for western Germany.

Of interest is the high value of ρ – the intra-class (establishment) correlation defined above – which at 0.93 indicates considerable inertia in collective bargaining status. In short, there is strong evidence that, after controlling for the X_i, the probabilities of an establishment being covered in any t_0 and t_1 are highly correlated. (The presence of non-trivial switching costs, referred to earlier, may of course lie at the root of this outcome.) Equivalently, the magnitude of σ_u (at 3.71) implies that a small difference in unobserved traits entails a quite different propensity of being covered by a collective agreement. Also note that since the significance test for ρ is itself a test for the presence of the unobserved effect ($\rho = 0$ if and only if $\sigma_u^2 = 0$), we can reject the simple pooled probit as an appropriate model description of the data.

(Table 2 near here)

The manifest intra-class correlation across distinct percentiles is given in Table 2.[5] Let us focus on the median percentile. Thus, setting the linear predictor at the median (the 0.50 column), the intra-class correlation is 0.76, flagging substantive within group persistence. Note also that for the median percentile, the corresponding joint probability in the second row (i.e. the probability of being covered in two given years) is equal to 0.47. In turn, the corresponding marginal probability of being covered by any type of agreement in any given year is 0.53 (first row), which closely tracks the (unweighted) mean coverage rate observed in the sample, at 52.7 percent.[6] Finally, the odds ratio in the fourth row indicates that the odds of an establishment being covered in t_0 and t_1 *versus* not being covered in t_0 but covered in t_1 are 56.2 times higher for the same observed characteristics. Since the odds ratio contrasts the (same) behaviour of two establishments in t_1, given that in t_0 they may have behaved differently, the conclusion is that it is considerably more likely that establishments that are covered will stay covered than non-covered establishments will join. Inertia in non-coverage is therefore very strong as well.

A glance at the full range of results presented in Table 2 reveals that the 0.25, 0.5 and 0.75 percentiles imply both a Pearson's *r* and an odds ratio that are very close to each other, while the marginal probability and the joint probability statistics exhibit a greater degree of variability. Interestingly, using Pearson's *r*, the 0.01 and the 0.99 percentiles show approximately the same degree of inertia in collective bargaining status, while the two very high odds ratios show that it is a considerably more likely that establishments in these two groups will stay covered in any two periods than join a collective agreement of any type.

Finally, by squaring the Pearson's *r* coefficient (at the median percentile), we obtain the result that collective bargaining coverage in a given year explains about 58 percent of the variation in collective bargaining behaviour in another year. The inference is that there is no terminal inertia in collective bargaining status, which result offers more than sufficient justification for an analysis of transitions into and out of collective bargaining.

IV. Transitions and Collective Bargaining 'Survivability'

We have seen that certain characteristics are associated with collective bargaining coverage. But can we say that the longer lasting its coverage, the less likely an establishment will be to change bargaining status? Our concern is now with the specific factors that induce *failure*, that is, transitions into or out of a collective agreement. The proper context for such analysis is survival modeling.

In our observation window, we have a maximum of nine annual observations which is insufficient to allow us to follow all production units from outset (birth) to death. The typical unit in our panel is indeed one that was born before 2000 and surveyed over a certain number of years within the observation interval. Figure 1 illustrates the array of possibilities. Establishment A, for example, was born before 2000 and is observed consecutively from 2000 up to point *e* (exit from a given state or point of 'failure').[7] Establishment A has therefore a left-truncation point as it is not possible to recover its bargaining status prior to 2000. Establishment B, on the other hand, is not only left-truncated but also right censored as well since it rotates out of the panel at point c. For their part, establishments C, D and E are observed for a number of years up to (a) ‘failure’, (b) attrition, and (c) right censoring (in 2008), respectively. Establishments F and G were born after 2000 and are, respectively, right censored and exiting a given state before 2008. Finally, there are those ‘permanent’ establishments, represented by case H, which are both left- and right-censored (in 2000 and 2008, respectively). In general, we will not be able to know the exact length of all spells because it is simply not possible to recover the ‘missing’ information. On the other hand, newly-founded establishments – and, to some extent, permanent establishments (i.e. panel stayers) – are a special case and they will be used to explain the survivability of collective bargaining. Again in the interests of expositional convenience, we focus on the aggregate category of collective agreements of any type.

(Figure 1 near here)

In the limit, the probability of failure, given by the hazard function, is constant and independent of any establishment attribute. This case is not particularly helpful in the present context since we believe that the selected covariates do have an impact on the hazard rate. Thus, we assume that leaving (or joining) a collective agreement of any type is a function of an observed set of time-constant (e.g. industry dummies) *and* time-varying (e.g. establishment size) covariates.[8]

Our hazard function belongs to the family of proportional hazards (PH) models

$$h(t;X) = k_1(X)k_2(t), \tag{2}$$

where k_1 and k_2 are the same functions for all establishments and X is the vector of the selected covariates (see, for example, Lancaster, 1990, Chapter 3). Setting $k_2(t) \equiv h_0(t)$, and $k_1(X) = \exp(X\beta)$, we have the standard Cox proportional hazards model

$$h(t;X) = h_0(t)\exp(X\beta), \tag{3}$$

where $h_0(t)$ is the baseline hazard (or the hazard rate when all covariates are set at zero).[9] Thus, $h(t)$ denotes, for covered (uncovered) establishments, the probability of an establishment leaving (joining) a collective agreement of any type, given that it has been covered (uncovered) up to time *t*. Given the longitudinal nature of our dataset, the standard errors of the estimated hazard coefficients are adjusted to account for the possible intra-group (establishment) correlation.

We also used parametric methods so that we can compare (a) PH estimates from both parametric (Weibull) and semiparametric (Cox) models, and (b) parametric estimates across a variety of models (Weibull, log-normal, and log-logistic).[10] Typically, and again omitting the subscript denoting the individual (establishment), a parametric model is written (in accelerated failure-time form) as $\log t = X\beta + z$, with $\log t$ denoting the logarithm of the survival time, and z the error term. z can assume different distributional forms, determining the regression model.

As mentioned earlier, we have both stock and flow sampling in our data, in the sense that we are able to observe *entrants* (newly-founded establishments) and non-entrants (i.e. establishments born at some point in the pre-observation period.[11] In the case of non-entrants, for whom left-censoring is the key problem, some further data manipulation will be required. For entrants, the survival analysis is straighforward since all spells for these units are either complete or right censored. In this context, the subsample of births turns out to be extremely useful, and we will discuss below the extent to which inferences based on births can be carried forward, first, to the subset of permanent establishments and then to the entire sample of surveyed units.

(Table 3 near here)

As shown in Table 3, we observe 2,679 births in the 1999-2006 period (plus 504 births in 2007 that are only observed once, namely in 2008). Of the total number of births, there are 266 collective agreement transitions in the 2001-2008 interval, comprising 149 leavers and 117 joiners. In other words, 9.9 percent of all births either switched into or out of a collective agreement during the sample period.

Table 3 also gives the collective agreement status in the year of birth *and* in the year of exit for all births in the sample, as well as the average year of exit (i.e. either rotation out of the panel due to attrition or transition into a different state) for each cohort. For example, an establishment born in 1999 is observed over an average period of 2.6 years before switching to a different regime or leaving the panel. Interestingly, the expected year of exit for our sample is virtually the same for covered and uncovered establishments. In any event, for establishments born later in the period, the average number of years prior to exit is necessarily smaller given that their number of years in the observation window becomes shorter.

From the total number of births in our dataset, and ignoring the 2007 cohort for which no transitions can be observed, in 52.2 percent of the cases establishments remain uncovered and 37.9 percent remain covered. This implies, as we have seen, that in 9.9 percent [100-(52.2+37.9)] of the cases we do observe establishments changing – either leaving or joining – their collective agreement status. Of those plants that are covered in the year of birth (plus one), some 12.8 percent do switch out of collective agreement within the observation window, while 7.7 percent of their non-covered counterparts will join a collective agreement. (Multiple failures – establishments with more than one transition over the observation period – are now dropped from our sample.)[12]

The results of model (2) – the hazard function – are presented in Table 4 for the two possible failure events: leaving a collective agreement or joining one. And in the last row of the table is given the median duration of "coverage/uncoverage", based on a PH exponential model without covariates. According to our estimates, the median duration of coverage for newly-founded establishments is 1.81 years, while the median duration without coverage is 2.61 years.

(Tables 4 and 4A near here)

As for the role of the selected covariates, greater establishment size decreases the probability of leaving a collective agreement, as does the use of modern technology. In contrast, foreign ownership and single-establishment status are associated with a higher failure rate. Note that the role of single-establishment status and foreign-owned variables are particularly strong; in particular, being a single establishment implies an 83 percent higher hazard rate, while foreign-ownership increases the hazard by 58 percent. In turn, a 1 percent increase in establishment size reduces the hazard by 0.35 percent. All other covariates included in the regression are not statistically significant.

The results for joining a collective agreement of any sort (given in the second column of Table 4) look symmetric. In particular, whereas the probability of leaving a collective agreement is found in the first column of the table to be decreasing in employment size, it is increasing in employment size when it comes to joining an agreement. But no other covariate is found to be statistically significant. *Vulgo*: establishment size is the major determinant of joining a collective agreement. That the evidence is much weaker in the case of transitions into collective agreements is not altogether unexpected given the smaller number of establishments engaging in such switching behaviour.

As shown in the penultimate row of Table 4, we do not find any evidence that the PH assumption has been violated, either in the case of leavers (first column of the table) or joiners (second column). For completeness, however, we also provide in Table 4A both the (PH) Weibull estimates and the non-PH parametric estimates from the log-normal and log-logistic models. First of all, observe that the median duration calculation from assuming a Weibull distribution rather than the exponential case is practically unchanged (cf. 1.81 in Table 4 with 1.89 in Table 4A in the case of leavers, and 2.61 in Table 4 with 2.58 in Table 4A for joiners). Second of all, by multiplying the Weibull coefficient in AFT form (see the first and second columns of Table 4A) by the shape parameter of the Weibull distribution, *p*, we obtain the corresponding Cox model coefficient. The comparison is quite striking in that (a) the sign of each coefficient is exactly the same, (b) their statistical significance is virtually unchanged, and (c) their magnitude is very close – striking but not surprising, as we did not reject the PH assumption in Table 4.

Table 4A also presents the results from two other parametric implementations, namely the log-normal and the log-logistic cases. Here we have to compare across the first, third and fifth columns, on the one hand, and across the second, fourth and sixth columns on the other. Again the results are very reassuring, in terms of the signs, statistical significance, and magnitudes of the estimated coefficients. Given these results, the Akaike information criterion (AIC), reported in the table, does not vary much across the selected alternatives, with the log-normal case giving the best fit (the smaller is the AIC value, the better).

It will be recalled that in our observation window all units are left-censored except for newly-founded plants. Since we cannot recoup the entire record of participation in collective bargaining for the former units, we can either ignore all transitions other than those observed for births or, alternatively, we can try to figure out an alternative procedure that avoids losing the valuable transition information we have on other types of establishments.

We elected to follow the second route, and therefore create a pre-observation period in which collective agreement status is unchanged for all units included in the risk analysis. To this end, we (a) divide the 2000-2008 period into the two sub-periods 2000-2003 and 2004-2008, (b) use the set of permanent establishments since these units are observed for a reasonably long time interval, and (c) impose the additional restriction of no change in status between 2000 and 2003. Transitions in the 2004-2008 interval will then be used to estimate the hazard. We will refer to this sample as the "restricted sample of permanent establishments". (Note that in enlarging the 'pre-observation' period from 2000-2003 to 2000-2004, for example, we reduced the risk period with no appreciable change in the results, other than a slight decrease in significance levels.)

In a second stage, and to test the role of left-censoring in our results – and ultimately evaluate whether the use of left-censored data in our survival analysis is legitimate – we added to the restricted sample of permanent establishments all those units in which collective bargaining status prior to 2004 is not fixed.[13] Taking, for example, the case of covered establishments this *counterfactual* exercise serves to compare the results from an experiment in which the left-censored units are necessarily covered with the case in which the presumed fixed coverage prior to 2004 is false for some units – and similarly for the case where the initial state is non-coverage. If the determinants of the hazard rate in the two counterfactual experiments are not too different (that is, where the hazard is not too sensitive to changes in the selected samples), we may conclude that left-censoring for permanent units of the panel is not really an issue, and that running the survival analysis on an *unrestricted* set of permanent establishments is not too much of a stretch. In this vein, our third and final exercise applies the survival model to all permanent establishments observed in 2000-2008 period, without further restrictions. Again, in this case we are simply ignoring left-censoring, implicitly assuming that either there was no change in status in the past (i.e. before 2000) or, alternatively, that it occurred too long ago to be a matter of concern.

We have 1,448 establishments in the restricted estimation sample of permanent establishments, of which 821 (627) were covered (uncovered) over 2000-2003. Of those that were covered (in 2000-2003), 93 subsequently switched out of collective bargaining between 2004 and 2008 – 93 out of 821 cases, or 11 percent. Of those that were not covered, 35 switched into collective agreements after 2003 – 35 out of 627 cases, or 6 percent.

The corresponding survival analysis, shown in the first column of Table 5, again indicates that establishment size is critical: the larger the establishment, the lower the probability that a covered establishment will leave a collective agreement. The single establishment variable is also statistically significant and positively signed as expected. The remaining variables are statistically insignificant. In turn, as shown in the second column of the table, joining collective agreements is less common among permanent stayers than among newly-founded establishments. Not surprisingly, therefore, all variables in the second column are statistically insignificant, with the sole exception of the establishment age dummy. In this case, older establishments tend to have a lower exit rate (from non-coverage). Apparently, non-covered establishments tend to stay non-covered, while the considerable minority that join collective agreements do not seem to share any readily discernible characteristics.

(Table 5 near here)

The second experiment – the *counterfactual* – is reproduced in Table 6. In this exercise, we added some 47 establishments to the sample in the first column of Table 5. The results are basically unchanged. The same obtains with respect to the transition behaviour of initially uncovered establishments, shown in the second column, where some 100 establishments have been added to the sample. The main lesson from the counterfactual is, then, that within the subsample of permanent establishments there seems to be no particular penalty in ignoring left-censoring.

(Tables 6 and 7 near here)

Given these findings, the final step is to present the survival analysis for the full set of permanent establishments. This procedure yields an enlarged estimation sample of 1,597 units, surveyed consecutively from 2000 to 2008. Of this total, there are 922 (675) establishments that were covered (not covered) by any type of collective agreement in 2000, and some 275 transitions comprising 193 leavers and 82 joiners. The results are presented in Table 7. As expected, the findings reported in the table mimic those obtained earlier in Table 5. From this perspective, it appears legitimate to conclude that in the case of permanent panel members there is enough evidence to support the proposition that plant size and skill content of the workforce matter in terms of collective bargaining survivability, while single establishment status favours the abandonment of collective bargaining. The influence of the remaining covariates on survivability of collective agreements is statistically weak, with the exception of the establishment age variable. However, it is more difficult to discern equally strong patterns in respect of transitions into collective agreements. Here, size and, to some limited extent, foreign ownership are the unique determinants, with again strikingly symmetric effects.

Finally, although the PH test rejects the null in only one case (viz. leaving any type of collective agreement, shown in the first column of Table 7), we also ran the corresponding parametric models for all cases contained in Tables 5 through 7. The results are not reported here but are available from the authors upon request. The principal findings are twofold. First, in both sign and statistical significance, the estimated coefficients are practically unchanged vis-à-vis those shown in Tables 5 through 7 – even in the case where the PH test rejected the null. Second, although the differences across the three parametric models are minor, the log-normal model offers the best fit.

V. Conclusions

We obtain confirmation in this study of the steady decline of sectoral bargaining in one of its hitherto more entrenched country settings. We can also confirm that the main source of the observed decline is attributable to (unexplained) establishment behaviour, holding establishment characteristics constant, rather than to any particularly unfavourable changes in observed characteristics.

More importantly, the analysis of the determinants of collective bargaining membership is in accordance with some of not all of our priors; that is, the skill content of the workforce, and the age and size of the establishment increase the propensity to be covered, while single-establishement status decreases it.

But we have also detected strong inertia in collective bargaining status, having controlled for establishment characteristics. This result is not unexpected given the long lasting presence of industry-wide agreeements in Germany. Specifically, we estimate that the odds of an establishment being covered in two different years, say t_0 and t_1, *versus* not being covered in t_0 but covered in t_1 are more than 50 times higher for the same observed characteristics.

Nevertheless, despite this observed persistence in collective agreements (and, equivalently, in their absence), establishments do switch their collective bargaining status. In this context, the finding that coverage (and absence of coverage) has a short half-life in the case of newly-founded establishments is of particular importance as it proves that collective bargaining is also volatile. Moreover, the finding that the estimated duration of coverage is shorter than the duration of no coverage, for this sample of establishments, also provides an interesting explanation for the observed decline in collective bargaining over the sample period.

Finally, the results from the survival analysis on the subset of left-censored permanent establishments show that the factors governing the decision to leave/join are after all general and not specific to any subset of establishments. Expressed another way, plant size and skill content matter in the decision to join, while single-plant establishments with outdated technology may be those opting out of collective agreeements.

The present treatment offers for the first time estimates of the median duration of coverage for newly-founded establishments. Further, using a set of counterfactual exercises, it is also shown that the determinants of collective bargaining transitions can be comfortably generalized to all types of establishments, both new and old. Our results are moreover robust to a variety of specifications. At root they suggest that, whatever position is taken on the current state of the German model, there is sufficient across-the-board variability in establishment behavior to indicate no lack of vitality.

Endnotes

1. In particular, we used a more restrictive definition of not-for-profit organizations while at the same time imputing collective bargaining status in all cases where one-year status changes likely reflected coding error (see the Data Appendix). A more restrictive definition of not-for-profit organizations means that we dropped not only public corporations from the sample but also those establishments that reported a "total budget" rather than sales volume. On net, these innovations resulted in a slightly larger sample; specifically, an additional 5,000 observations.

2. A different choice of reference groups would yield $(x_{2008} - x_{2000})b_{2008}$ and $x_{2000}(b_{2008} - b_{2000})$ for the between and within effects, respectively. Our results are not sensitive to the choice of the reference groups.

3. Unfortunately, not imposing these restrictions on the relationship between u_i and X_{it}, in the spirit of the fixed-effects analysis, is infeasible due to the *incidental parameters problem* (Wooldridge, 2002: 484). We cannot therefore offer any fixed-effects versus random-effects test. As it will be shown below, we can easily reject the null of no presence of the unobserved effect, u_i. This finding is the reason why we do not provide any results based on a simple pooled probit specification. Finally, relaxing the random-effects model assumptions is a complex issue that is beyond the scope of the present paper the main goal of which is to provide indicative estimates of the propensity to be covered by a collective agreement.

4. The equivalent latent variable model is given by $Y_{it}^* = X_{it}\beta + u_i + e_{it}$ where Y_{it}^* is the latent variable and $e_{it} \sim N(0, 1)$, with e_{it} uncorrelated with u_i. Assuming $\Pr(Y_{it} = 1|u_i, X_{it}) = \Pr(Y_{it}^* > 0|u_i, X_{it})$, model (1) follows readily.

5. We cannot perform a similar exercise for ρ since it does not depend on the marginal distribution.

6. 52 percent is the unweighted average coverage rate in the estimation sample. According to Table A.3, the weighted average is necessarily lower than 51.2 percent as the coverage rate is roughly monotonically decreasing over the 2000-2008 period. The difference between weigthed and non-weigthed averages has to do with the well-known over-representation of large firms in the IAB survey.

7. We note that by 'failure' we mean an exit from a given collective bargaining state, not an exit from the market (death). All establishment deaths were excluded from the panel.

8. For the time-varying covariates, we shall ignore possible anticipation and delay effects. We shall also assume that the effect of any continuous variable on the hazard is independent of the level of the variable (i.e. the marginal effect is constant). A model without covariates will be used to obtain the predicted median duration of coverage/absence of coverage for newly-founded establishments (see last row of Table 4).

9. Formally, the model in equation (3) is PH with time-invariant covariates. The corresponding PH model with time-varying variables is given by $h(t; X(t)) = h_0(t)\exp(X\beta)$ (see Wooldridge, 2002: 693).

10. The generalized gamma model did not achieve convergence in our implementation. Since the number of transitions is limited in our dataset and the set of selected covariates includes a number of dummy variables, this result is not overly surprising. We eschewed reducing the set of regressors in order to achieve convergence of the generalized gamma given the need to work a uniform set of covariates throughout the paper.

11. The year of birth of any establishment in the panel is always known. Only the bargaining status in the pre-observation period is unknown.

12. The number of multiple failures in the sample of births is 195, out of 2,874 cases (or 6.8 percent). Of all these multiple failures, 76 percent record more than two switches, which we regard as indicative of bad measurement. Accordingly, all multiple failures were excluded from our regressions. We should further note that transitions are even less frequent among permanent establishments (cf. Tables 5 through 7) and therefore for this subsample multiple collective bargaining changes are a matter of still less concern.

13. For transitions into collective agreements, this amounts to adding the following sequences to the existing restricted sample of permanents: 0111|11111, 0011|11111, 0001|11111, and 0000|11111. In the case of transitions out of collective agreements, we add the sequences 1111|00000, 1110|00000, 1100|00000, and 1000|00000. The vertical bar in these sequences denotes the 2003 separation point and '1' ('0') signifies coverage (absence of coverage or 'uncoverage'). The 2004-2008 interval defines the risk period.

References

Addison J.T., Bryson A., Teixeira P. and Pahnke A. (2011) 'Slip Sliding Away: Further Union Decline in Germany and Britain', *Scottish Journal of Political Economy*, 58: 490–518.

Addison J.T., Bryson A., Teixeira P., Pahnke A. and Bellmann L. (2010) 'The State of Collective Bargaining and Worker Representation in Germany: The Erosion Continues', IZA Discussion Paper No. 5053, Institute for the Study of Labor, Bonn, Germany.

Bispinck R., Dribbusch H. and Schulten, T. (2010) 'German Collective Bargaining in a European Perspective. Continuous Erosion or Re-Stabilization of Multi-Employer Agreements? ', Wirtschafts- und Sozialwissenschaftliches Institut, Düsseldorf: Hans Böckler Stiftung.

Calmfors L. and Driffill, J. (1988) 'Bargaining Structure, Corporatism and Macroeconomic Performance', *Economic Policy*, 3: 13–61.

Doellgast V. and Greer I. (2007) 'Vertical Disintegration and the Disorganization of German Industrial Relations', *British Journal of Industrial Relations*, 45: 55–76.

Dustman C., Ludsteck J. and Schönberg U. (2009) 'Revisiting the German Wage Structure', *Quarterly Journal of Economics*, 124: 843–881.

Ellguth P. and Kohaut S. (2010) 'Tarifbindung und betriebliche Interessenvertretung – Aktuelle Ergebnisse aus dem IAB-Betriebspanel 2009', *WSI-Mitteilungen*, 63: 204–209.

Ellguth P. and Kohaut, S. 2011 'Tarifbindung und tarifliche Interessenvertretung: Aktuelle Ergebnisse aus dem IAB-Betriebspanel', *WSI-Mitteilungen*, 64: 242–247.

Fischer G., Janik F., Müller D. and Schmucker A. (2009) 'The IAB Establishment Panel: Things Users Should Know', *Schmollers Jahrbuch*, 129: 133–148.

Frege C. M. (2003) 'Transforming German Workplace Relations: Quo Vadis Cooperation?', *Economic and Industrial Democracy*, 24: 317–347.

Haipeter T. and Lehndorff S. (2009) 'Collective Bargaining on Employnment', Working Paper No. 3, Industrial Relations Department, Geneva: International Labour Office.

Hassel A. (1999) 'The Erosion of the German System of Industrial Relations', *British Journal of Industrial Relations*, 37: 483–505.

Hassel A. (2002) 'The Erosion Continues: Reply [to Klikauer]', *British Journal of Industrial Relations*, 40: 309–317.

Klikauer T. (2002) 'Stability in Germany's Industrial Relations: A Critique on Hassel's Erosion thesis', *British Journal of Industrial Relations*, 40: 295–308.

Kohaut S. and Schnabel C. (2007) 'Tarifliche Öffnungsklauseln: Verbreitung, Inanspruchnahme und Bedeutung', *Sozialer Fortschritt*, 56: 33–40.

Kohaut S. and Schnabel C. (2003a) 'Tarifverträge – nein danke? Ausmaß und Einflussfaktoren der Tarifbindung west- und ostdeutscher Betriebe', *Jahrbücher für Nationalalökonomie und Statistik*, 223: 312–33.

Kohaut S. and Schnabel C. (2003b) 'Zur Erosion des Flächentarifvertrags: Ausmaß, Einflussfaktoren und Gegenmaßnahmen', *Industrielle Beziehungen*, 10: 193–219.

Lancaster T. (1990) *The Econometric Analysis of Transition Data*, Cambridge: Cambridge University Press.

OECD (2004) *Employment Outlook,* Paris: Organisation for Economic Co-operation and Development.

Rodríguez G. and Elo I. (2003) 'Intra-class Correlation in Random-Effects Models for Binary Data', *The Stata Journal*, 3: 32–46.

Schnabel C., Zagelmeyer S. and Kohaut S. (2006) 'Collective Bargaining Structure and its Determinants: An Empirical Analysis with British and German Establishment Data', *European Journal of Industrial Relations*, 12: 165–88.

Seifert H. and Massa-Wirth H. (2005) 'Pacts for Employment and Competitiveness in Germany', *Industrial Relations Journal*, 36: 217–240.

Willman P., Bryson A. and Gomez R. (2007) 'The Long Goodbye: New Establishments and the Fall of Union Voice in Britain', *International Journal of Human Resource Management*, 18: 1318–1334.

Wooldridge J. (2002) *Econometric Analysis of Cross-Section and Panel Data*, Cambridge. MA: The MIT Press.

Data Appendix

Our data is extracted from the IAB Establishment Panel (or *Betriebspanel*), which is based on a stratified random sample of plants from the population of all establishments with at least one employee covered by social insurance. A full description of the *Betriebspanel* is provided by Fischer *et al.* (2009). We confine our remarks here to outlining the procedures used to generate our various estimation samples.

Firstly, we took the 2008 survey and appended to it all previous surveys back to 2000. We decided not to range further back in time primarily to avoid having to deal with changes in industry classification in 2000 (from a 3- to a 5-digit system), while extending the sample beyond 2008 would also require further manipulation of the data as industries were again reclassified in 2009. Secondly, we focus on establishments from the private, profit-oriented sector of the economy with at least 5 employees. Smaller establishments were excised on pragmatic grounds: they typically evince more spurious volatility in collective bargaining coverage. In total, we have 9 surveys (from 2000 through 2008), encompassing some 82,000 observations on approximately 24,000 establishments in the whole of Germany. We note that our criteria for excluding not-for-profit organizations implied that all public-sector units were excluded from the sample, as well as all those institutions that failed to report some sales volume.

The constructed longitudinal dataset (i.e. the panel in which establishments are followed over time for a maximum period of nine years) is used to analyze collective bargaining membership, on the one hand, and the duration of collective agreements, on the other. Note that in general we do not know the elapsed duration of the observed spells. That is to say, we do not know the number of years in which a given establishment has been either covered or uncovered at the point it is first observed in the survey. All establishments are therefore left-truncated, with the notable exception of the newly-founded establishments (i.e. births) that we were able to follow from the outset. To ensure that the year of birth coded in the survey panel was correct, we used the establishment register (or *Betriebsdatei*) and the fact that establishments in the two raw datasets (i.e. *Betriebsdatei* and *Betriebspanel*) share exactly the same identification code (or *Betriebsnummer*). Finally, in coding the collective agreement variable, we assumed that if the status in year $t-1$ was the same as in year $t+1$, then the status in year t was unchanged. This assumption resulted in 3.3 percent of all collective agreement observations being recoded.

The selected covariates are presented in the table below. They comprise two measures of workforce composition based on skill and gender, foreign ownership, single establishment versus multi-site firm status, establishment age, establishment size, and an indicator of the state of technology in use. The latter is a dummy equal to 1 if the technology of equipment is high or very high, 0 otherwise. (Note that since the technology question was not asked in 2004 we used the mean of 2003 and 2004 scores to complete the series.) These arguments are augmented by a total of thirty seven 2-digit industry dummies plus sixteen regional dummies. Although somewhat sparse, our choice of regressors is guided by the literature *and* the need to minimize the loss of establishments occasioned by missing observations, especially in a situation where the number of collective agreement transitions is limited.

Variable		Full sample		Births	
		Mean	*n*	Mean	*n*
Any type of collective agreement	Dummy	0.527	82,137	0.435	8,158
Sectoral agreement	Dummy	0.458	82,137	0.369	8,158
Firm-level agreement	Dummy	0.069	82,137	0.067	8,158
Log number of employees	Continuous	3.685	82,137	3.240	8,158
Use of modern technology	Dummy	0.693	80,146	0.707	7,641
Proportion of skilled workers	Percent	67.36	82,118	65.13	8,156
Proportion of female workers	Percent	37.85	82,004	39.88	8,150
Foreign majority ownership	Dummy	0.072	80,715	0.079	8,027
Single establishment	Dummy	0.713	81,400	0.692	8,087
Establishment older than 10 years	Dummy	0.650	81,769		
Regional dummies (16)	Dummy		82,137		8,158
Industry dummies (37)	Dummy		82,137		8,158

TABLE 1
Coverage Propensity by Type of Collective Agreement, Random-Effects Probit Estimates, 2000-2008.

	Germany	West	East
Log number of employees	0.977*** (0.021)	1.008*** (0.029)	0.925*** (0.034)
Use of modern technology	0.019 (0.030)	0.010 (0.040)	0.037 (0.046)
Proportion of skilled workers	0.004*** (0.001)	0.004*** (0.0008)	0.004*** (0.001)
Proportion of female workers	-0.001 (0.001)	-0.002* (0.001)	0.001 (0.002)
Foreign majority ownership	0.155* (0.085)	0.018 (0.102)	0.452*** (0.156)
Single establishment	-0.643*** (0.045)	-0.480*** (0.058)	-0.883*** (0.074)
Establishment older than 10 years	1.176*** (0.068)	1.127*** (0.091)	1.073*** (0.097)
Regional dummies	Yes	Yes	Yes
Industry dummies	Yes	Yes	Yes
σ_u	3.714 (0.051)	3.858 (0.067)	3.389 (0.077)
ρ	0.932 (0.002)	0.937 (0.002)	0.920 (0.003)
Wald χ^2	7595.08	4264.23	2546.80
Number of observations	80,958	50,895	30,063
Number of establishments	24,018	16,187	7,835

Notes: The model is given by equation (1) in the text. σ_u is the standard deviation of the unobserved effect u_i, and ρ is the latent intra-group (establishment) correlation. The model specification in the first column also contains 16 regional dummies (10 and 6 for western and eastern Germany in the second and third columns of the table, respectively), 37 two-digit industry dummies, and 8 year dummies. Standard errors are given in parentheses.
***, **, * denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

TABLE 2

Marginal and Joint Coverage Probabilities and Intra-class Manifest Correlation

Germany	Percentiles				
	0.01	0.25	0.50	0.75	0.99
Marginal probability	0.047	0.340	0.526	0.715	0.965
Joint probability	0.032	0.286	0.467	0.665	0.954
Pearson's r	0.681	0.759	0.764	0.755	0.668
Odds ratio	152.118	59.430	56.248	62.534	182.264
West					
Marginal probability	0.049	0.435	0.627	0.789	0.973
Joint probability	0.035	0.379	0.573	0.748	0.965
Pearson's r	0.694	0.772	0.770	0.754	0.668
Odds ratio	160.996	61.485	63.083	75.693	235.672
East					
Marginal probability	0.059	0.229	0.369	0.538	0.895
Joint probability	0.040	0.180	0.308	0.474	0.866
Pearson's r	0.665	0.725	0.740	0.743	0.692
Odds ratio	107.931	55.406	47.896	46.309	78.754

Notes: The reported statistics are obtained using the *xtrho* command in Stata 10, and are described in Rodriguez and Elo (2003). The 95% confidence intervals for the median percentile and for Germany, for example, are (0.526, 0.527), (0.467, 0.468), (0.758, 0.770), and (52.681, 59.098), respectively. See section III for definitions.

TABLE 3
Beginning- and End-period Collective Agreement Status of Newly-founded Establishments, 2000-2008, Unweighted Data

Year of birth	Collective agreement status in year of birth+1		Collective agreement status in year of exit		Year of exit (average)
	Status	*n*	*Anycb*=0	*Anycb*=1	
1999	*Anycb*=0	150	132	18	2002.6
	Anycb=1	124	19	105	2002.6
	Total	274	151	123	2002.6
2000	*Anycb*=0	138	129	9	2003.1
	Anycb=1	118	17	101	2003.4
	Total	256	146	110	2003.2
2001	*Anycb*=0	172	159	13	2004.3
	Anycb=1	112	11	101	2004.1
	Total	284	170	114	2004.2
2002	*Anycb*=0	68	64	4	2005.1
	Anycb=1	38	6	32	2005.9
	Total	106	70	36	2005.4
2003	*Anycb*=0	253	231	22	2006.0
	Anycb=1	198	20	178	2006.0
	Total	451	251	200	2006.0
2004	*Anycb*=0	203	185	18	2006.7
	Anycb=1	195	37	158	2006.7
	Total	398	222	176	2006.7
2005	*Anycb*=0	241	230	11	2007.3
	Anycb=1	178	17	161	2007.3
	Total	419	247	172	2007.3
2006	*Anycb*=0	290	268	22	2007.7
	Anycb=1	201	22	179	2007.7
	Total	491	290	201	2007.7
2007	*Anycb*=0	278			
	Anycb=1	226			
	Total	504			

Notes: A newly-founded establishment in the 2000 (2001, …, 2008) survey is a unit born in 1999 (2000, …, 2007). Consequently, all year 2008 births (i.e. establishments born in 2008) are discarded in our subsequent survival analysis. Also note that all establishments born in, say, 2002 but not observed (surveyed) before 2006, for example, are dropped from the sample. In other words, only those establishments that can be followed from the outset (year of birth) are included in the estimation sample. Exit means rotation out of the panel due either to attrition or failure (end of the initial state). *Anycb* is a dummy variable signifying the presence of any type of agreement.

TABLE 4
Cox Proportional Hazards Model Estimates, Newly-founded Establishments, 2000-2008

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.348 (0.068)***	0.349 (0.092)***
Use of modern technology	-0.500 (0.157)***	0.011 (0.203)
Proportion of skilled workers	-0.004 (0.003)	0.002 (0.004)
Proportion of female workers	-0.007 (0.004)*	0.001 (0.004)
Foreign majority ownership	0.460 (0.273)*	-0.490 (0.449)
Single establishment	0.604 (0.215)***	-0.032 (0.245)
Number of observations	1,787	2,362
Number of establishments	787	1,003
Number of failures	145	117
Wald χ^2	81.47	73.91
PH test (χ^2)	12.07	15.24
Predicted median duration	1.81	2.61

Notes: The hazard function is given by equation (2). The model includes 7 industry dummies and 1 regional dummy (western Germany). Clustered standard errors are given in parentheses. The Wald test rejects the null of no joint statistical significance of the model at the .01 level. The PH test, based on the Schoenfeld residuals, examines whether the proportional hazards (PH) assumption holds for the selected set of covariates. In neither column is there evidence that the PH assumption has been violated. The (predicted) median duration in the last row of the table is obtained using a PH exponential model without covariates.

TABLE 4A

Parametric Model Estimates, Newly-founded Establishments, 2000-2008

	Weibull		Log-normal		Log-logistic	
	Leaving	Joining	Leaving	Joining	Leaving	Joining
Log number of employees	0.256*** (0.050)	-0.211*** (0.058)	0.206*** (0.040)	-0.156*** (0.049)	0.229*** (0.045)	-0.195*** (0.053)
Use of modern technology	0.420*** (0.117)	-0.009 (0.124)	0.348*** (0.096)	-0.003 (0.106)	0.394*** (0.105)	0.006 (0.118)
Proportion of skilled workers	0.004* (0.002)	-0.0006 (0.002)	0.004** (0.002)	-0.0006 (0.002)	0.004** (0.002)	-0.0009 (0.002)
Proportion of female workers	0.006* (0.003)	-0.0005 (0.003)	0.004* (0.002)	-0.0008 (0.002)	0.004* (0.002)	-0.0006 (0.002)
Foreign majority ownership	-0.337 (0.211)	0.338 (0.279)	-0.197 (0.171)	0.310 (0.219)	-0.252 (0.177)	0.336 (0.256)
Single establishment	-0.465*** (0.165)	0.072 (0.154)	-0.347*** (0.118)	0.138 (0.125)	-0.406*** (0.140)	0.119 (0.139)
Number of observations	1,787	2,362	1,787	2,362	1,787	2,362
Number of establishments	787	1,003	787	1,003	787	1,003
Number of failures	145	117	145	117	145	117
Wald χ^2	67.01	71.18	82.06	77.83	92.03	85.95
AIC	701.58	651.29	691.16	642.51	695.32	645.15
Predicted median duration	1.89	2.58	1.66	2.22	1.45	2.01

Notes: These reported results are all in accelerated failure-time (AFT) form. In the Weibull case, to obtain a coefficient comparable with that generated by the Cox model one has simply to multiply the reported value by (-p), where p is the shape parameter of the Weibull distribution. The results from the two models are directly comparable in their magnitude, sign, and statistical significance. (p is equal to 1.435501 (1.693232) in the first (second) column of the table). AIC denotes the Akaike information criterion. The (predicted) median duration given in the last row of the table is obtained using the corresponding distribution in a model without covariates.

TABLE 5
Cox Proportional Hazards Model Estimates, Restricted Sample of Permanent Establishments, 2004-2008

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.241 (0.074)***	0.050 (0.237)
Use of modern technology	0.150 (0.232)	0.447 (0.447)
Proportion of skilled workers	-0.002 (0.005)	0.007 (0.012)
Proportion of female workers	-0.001 (0.005)	-0.001 (0.009)
Foreign majority ownership	-0.788 (0.598)	0.434 (0.855)
Single establishment	1.002 (0.303)***	-0.431 (0.486)
Establishment age	0.072 (0.280)	-0.694 (0.385)*
Number of observations	3,928	3,051
Number of establishments	821	627
Number of failures	93	35
Wald χ^2	76.89	8,783.72
PH test (χ^2)	12.96	11.82

Note: See notes to Table 4.

TABLE 6
Cox Proportional Hazards Model Estimates, Restricted Sample of Permanent Establishments, 2004-2008, *Counterfactual*

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.224 (0.074)***	0.055 (0.236)
Use of modern technology	0.175 (0.234)	0.461 (0.445)
Proportion of skilled workers	-0.002 (0.005)	0.008 (0.012)
Proportion of female workers	-0.009 (0.005)*	-0.0001 (0.009)
Foreign majority ownership	-0.810 (0.597)	0.492 (0.853)
Single establishment	0.976 (0.305)***	-0.384 (0.478)
Establishment age	0.123 (0.287)	-0.763 (0.386)**
Number of observations	4,163	3,551
Number of establishments	868	727
Number of failures	93	35
Wald χ^2	75.45	28.22
PH test (χ^2)	13.08	11.97

Notes: See notes to Table 4.The Wald test in the second column of the table does not reject the null of no joint statistical significance of the model.

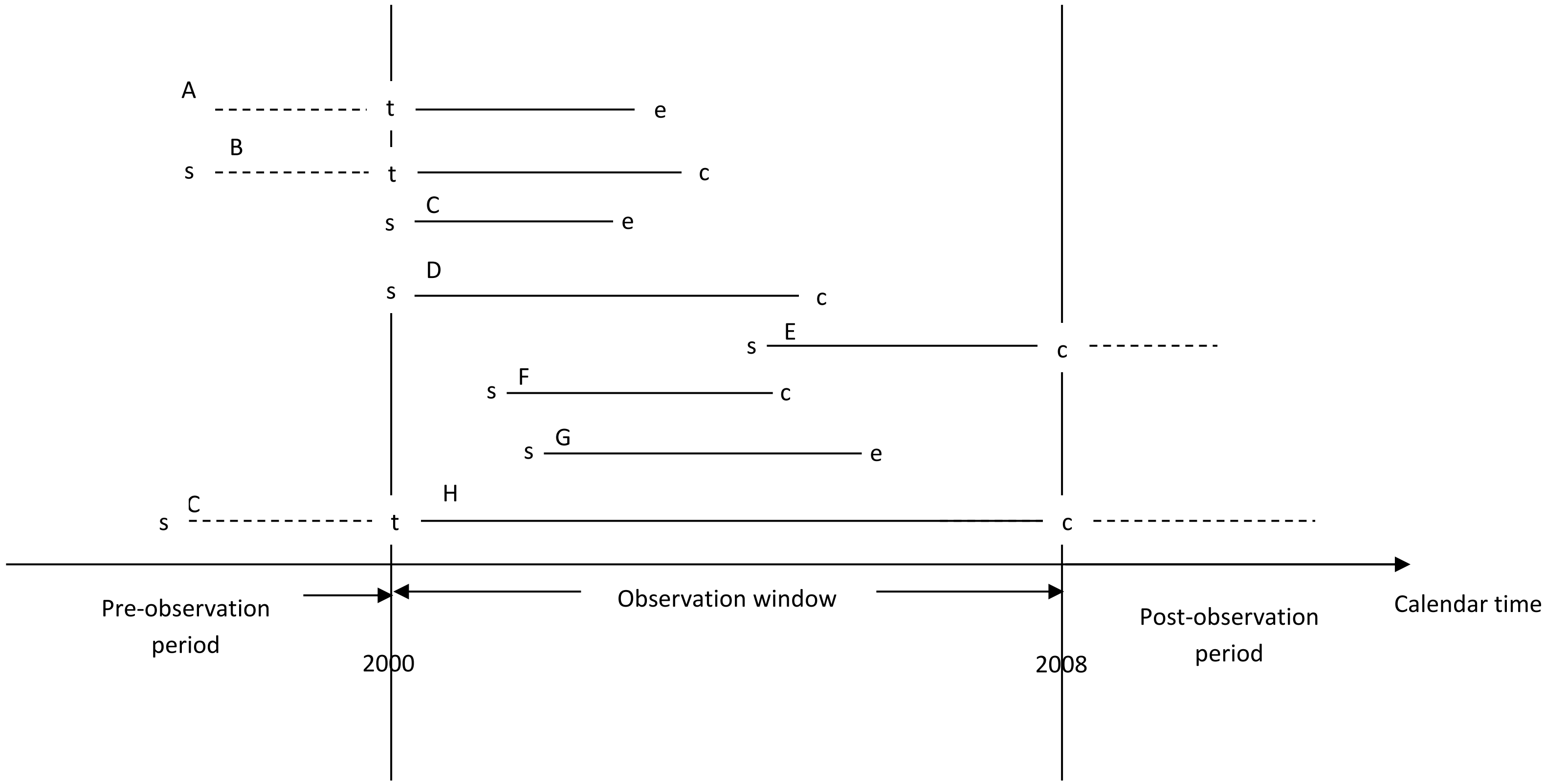
TABLE 7
Cox Proportional Hazards Model Estimates, Sample of Permanent Establishments, 2000-2008

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.367 (0.054)***	0.224 (0.122)**
Use of modern technology	0.245 (0.165)	0.193 (0.272)
Proportion of skilled workers	-0.006 (0.003)**	0.0004 (0.005)
Proportion of female workers	-0.005 (0.003)	-0.011 (0.006)**
Foreign majority ownership	-0.630 (0.422)	0.623 (0.462)
Single establishment	0.648 (0.198)***	-0.337 (0.298)
Establishment age	-0.212 (0.169)	0.074 (0.247)
Number of observations	7,486	5,697
Number of establishments	922	675
Number of failures	193	82
Wald χ^2	147.56	31.45
PH test (χ^2)	32.25	13.43

Note: See notes to Table 4. In the first column, the null hypothesis that the PH assumption holds is rejected.

FIGURE 1

Schematic of the observation window and censoring



Legend: t – left-truncation point
c – right-censoring point
s – starting time of the event (or entry to a state)
e – ending time of the event (or exit from a state)

TABLE A.1
Collective Bargaining Coverage by Employment and by Establishment, 2000-2008 (establishments with at least 5 employees, cross-section weighted data, percentage values)

		No Collective Agreement		Firm-level Agreement		Sectoral Agreement	
Year	Region	Employment	Establishment	Employment	Establishment	Employment	Establishment
2000	Germany	34.3	48.9	7.0	2.8	58.7	48.3
	West	31.5	44.4	6.4	2.4	62.1	53.2
	East	48.4	67.2	9.9	4.5	41.7	28.3
2001	Germany	33.7	49.9	7.9	3.0	58.4	47.1
	West	30.8	45.7	6.9	2.4	62.4	52.0
	East	47.9	67.3	13.2	5.7	38.9	27.0
2002	Germany	34.7	51.5	7.6	2.6	57.7	45.8
	West	31.7	47.0	6.7	2.1	61.5	50.8
	East	50.2	70.8	11.9	4.8	37.9	24.4
2003	Germany	35.2	52.8	7.7	2.3	57.1	44.8
	West	32.1	48.1	7.1	1.9	60.8	50.0
	East	50.8	72.3	10.7	4.2	38.5	23.5
2004	Germany	36.9	56.0	7.6	2.4	55.5	41.6
	West	33.9	52.0	7.0	2.1	59.2	45.9
	East	52.3	72.9	10.8	3.7	36.9	23.4
2005	Germany	38.6	57.0	7.8	2.5	53.5	40.5
	West	36.1	53.6	7.2	2.3	56.7	44.1
	East	28.4	71.3	5.9	4.2	65.6	24.5
2006	Germany	40.3	58.7	8.0	2.3	51.7	38.9
	West	38.0	55.8	7.1	1.9	54.8	42.4
	East	52.2	71.8	12.2	4.5	35.6	23.7
2007	Germany	41.7	59.9	7.5	2.7	50.8	37.4
	West	39.5	57.0	6.6	2.1	53.9	40.9
	East	53.5	72.8	11.8	5.1	34.8	22.1
2008	Germany	42.8	60.8	8.0	2.8	49.2	36.4
	West	40.6	58.1	7.3	2.3	52.1	39.5
	East	54.4	71.9	11.3	4.5	34.3	23.6

TABLE A.2
Collective Bargaining Coverage by Employment and by Establishment, 2000-2008, for Plants of Different Sizes (cross-section weighted data, percentage values)

Germany		No Collective Agreement		Firm-level Agreement		Sectoral Agreement	
Year	Establishment size	Employment	Establishment	Employment	Establishment	Employment	Establishment
2000	With at least 250	11.2	17.0	12.8	12.3	76.0	70.7
	With less than 250	43.2	49.3	4.7	2.7	52.1	48.0
2001	With at least 250	10.7	15.5	15.9	14.0	73.4	70.5
	With less than 250	42.9	50.3	4.7	2.9	52.3	46.8
2002	With at least 250	10.8	15.9	14.5	13.5	74.6	70.6
	With less than 250	44.0	51.9	4.9	2.5	51.1	45.6
2003	With at least 250	11.7	16.6	15.6	13.1	72.7	70.3
	With less than 250	44.4	53.2	4.6	2.2	51.0	44.6
2004	With at least 250	11.0	16.2	14.0	14.1	75.0	69.8
	With less than 250	45.4	56.4	5.3	2.3	49.3	41.3
2005	With at least 250	12.2	17.2	14.0	14.0	73.8	68.9
	With less than 250	48.3	57.3	5.5	2.6	46.2	40.2
2006	With at least 250	13.3	20.3	15.8	14.3	70.8	65.4
	With less than 250	50.2	59.1	5.1	2.2	44.7	38.6
2007	With at least 250	15.3	21.6	13.2	13.1	71.5	65.3
	With less than 250	51.6	60.3	5.3	2.6	43.1	37.1
2008	With at least 250	17.7	24.7	14.7	13.1	67.6	62.2
	With less than 250	52.3	61.2	5.5	2.6	42.2	36.2
West							
2000	With at least 250	10.7	16.2	12.0	11.2	77.3	72.6
	With less than 250	40.4	44.8	4.0	2.2	55.6	53.0
2001	With at least 250	10.2	15.1	14.1	11.8	75.7	73.1
	With less than 250	39.7	46.0	3.7	2.3	56.6	51.7
2002	With at least 250	10.2	15.4	13.2	11.7	76.6	72.8
	With less than 250	40.8	47.4	4.0	2.0	55.2	50.6
2003	With at least 250	11.0	16.0	14.8	11.6	74.2	72.5
	With less than 250	41.1	48.5	3.8	1.8	55.1	49.7
2004	With at least 250	10.2	15.3	13.0	12.9	76.8	71.9
	With less than 250	43.6	52.4	4.5	2.0	51.9	45.6
2005	With at least 250	11.6	16.6	13.2	12.7	75.2	70.7
	With less than 250	45.8	53.9	4.9	2.2	49.4	43.8
2006	With at least 250	12.5	19.0	14.6	13.2	72.9	67.8
	With less than 250	48.1	56.2	4.2	1.7	47.7	42.1
2007	With at least 250	14.7	20.8	12.0	11.4	73.3	67.7
	With less than 250	49.6	57.4	4.5	2.0	45.9	40.6
2008	With at least 250	16.9	24.0	13.5	11.2	69.6	64.8
	With less than 250	50.3	58.5	4.8	2.2	44.9	39.2

(cont.)

East		No Collective Agreement		Firm-level Agreement		Sectoral Agreement	
Year	Establishment size	Employment	Establishment	Employment	Establishment	Employment	Establishment
2000	With at least 250	15.8	23.2	20.3	20.3	63.9	56.5
	With less than 250	55.0	67.5	7.8	4.4	37.2	28.1
2001	With at least 250	13.8	18.1	29.4	27.3	56.8	54.6
	With less than 250	56.6	67.6	9.0	5.5	34.4	26.8
2002	With at least 250	15.6	19.0	25.1	24.7	59.3	56.3
	With less than 250	58.4	71.2	8.8	4.6	32.8	24.2
2003	With at least 250	17.0	20.6	21.5	22.3	61.5	57.1
	With less than 250	59.0	72.7	8.1	4.1	32.9	23.2
2004	With at least 250	17.7	21.8	21.9	21.8	60.4	56.4
	With less than 250	60.3	73.3	8.2	3.5	31.4	23.2
2005	With at least 250	16.9	20.8	21.4	22.4	61.6	56.8
	With less than 250	59.5	71.7	8.6	4.1	31.8	24.3
2006	With at least 250	20.0	28.1	25.1	20.8	54.9	51.0
	With less than 250	59.7	72.1	9.2	4.4	31.2	23.5
2007	With at least 250	20.6	26.5	22.7	23.3	56.6	50.2
	With less than 250	61.0	73.2	9.3	5.0	29.8	21.9
2008	With at least 250	24.4	29.5	24.3	26.0	51.3	44.5
	With less than 250	60.8	72.2	8.5	4.4	30.7	23.5

TABLE A.3

Within versus Compositional Change by Type of Agreement and by Region, 2000 and 2008, 2000 and 2004, and 2005 and 2008 (cross-section weighted data)

		Germany		West		East	
	(a) Collective agreements of any type, 2000-2008	2000	2008	2000	2008	2000	2008
(1)	Observed coverage rate	51.2	39.2	55.7	41.9	33.1	28.1
(2)	Percentage point change, 2000-2008		-12.0		-13.8		-5.0
(3)	2008 (predicted) coverage based on 2000 coefficients		51.0		55.5		34.1
(4)	2000 (predicted) coverage based on 2008 coefficients	40.0		42.7		28.1	
(5)	Percentage point change due to changes in characteristics based on 2000 coefficients		-0.2 (1.5%)		-0.2 (1.3%)		1.1 (-21.9%)
(6)	Percentage point change due to changes in behavior based on 2000 coefficients		-11.9 (98.5%)		-13.7 (98.7%)		-6.1 (121.9%)
	(b) Collective agreements of any type, 2000-2004	2000	2004	2000	2004	2000	2004
(1)	Observed coverage rate	51.2	45.0	55.7	49.1	33.1	27.6
(2)	Percentage point change, 2000-2004		-6.2		-6.6		-5.5
(3)	2004 (predicted) coverage based on 2000 coefficients		50.6		55.3		32.8
(4)	2000 (predicted) coverage based on 2004 coefficients	45.4		49.3		28.4	
(5)	Percentage point change due to changes in characteristics based on 2000 coefficients		-0.6 (9.3%)		-0.4 (6.1%)		-0.3 (5.5%)
(6)	Percentage point change due to changes in behavior based on 2000 coefficients		-5.6 (90.7%)		-6.2 (93,9%)		-5.2 (94.5%)
	(c) Collective agreements of any type, 2005-2008	2005	2008	2005	2008	2005	2008
(1)	Observed coverage rate	43.0	39.2	46.3	41.9	28.9	28.1
(2)	Percentage point change, 2005-2008		-3.8		-4,4		-0.8
(3)	2008 (predicted) coverage based on 2005 coefficients		43.8		47.2		29.7
(4)	2005 (predicted) coverage based on 2008 coefficients	38.6		41.1		27.8	
(5)	Percentage point change due to changes in characteristics based on 2005 coefficients		0.8 (-21.1%)		0.9 (-20.4%)		0.8 (-100.0%)
(6)	Percentage point change due to changes in behavior based on 2005 coefficients		-4.6 (121.1%)		-5.3 (120.4%)		-1.6 (200.0%)

Notes: The within effect is always statistically significant at the .01 level, while the between effect is never statistically significant. Taking the panel (a) results for illustrative purposes, the between effect in row (5) is given by row (3) minus row (1) for 2000, and the within effect in row (6) is given by row (2) minus row (5), or the 2000-2008 percentage point change in row (2) minus the between effect. Alternatively put, if $x_{2008}b_{2000}$ is the 2008 (predicted) coverage based on year 2000 coefficients, then the between effect is given by $(x_{2008}-x_{2000})\, b_{2000}$ and the within effect by $x_{2008}(b_{2008} - b_{2000})$, where the reference groups are the year 2000 coefficients and the year 2008 characteristics, respectively (see text). The results in panels (b) and (c) can be similarly interpreted. Our selected vector of covariates x includes establishment size, the proportion of skilled and female workers, and dummies for single-establishment status, foreign ownership, establishment age, state of technology, industry and region.